Credit-risk valuation in the sovereign CDS and bonds markets: Evidence from the euro area crisis

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Documentos de Trabajo
Nº 53

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Febrero/February 2012

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ISSN (edición impresa): 2172-6337

ISSN (edición electrónica): 2172-7147

Depósito Legal: BI-2910-2010

Maqueta e imprime: Composiciones Rali, S.A.

Abstract

We analyse the extent to which prices in the sovereign credit default swap (CDS) and bond markets reflect the same information on credit risk in the context of the European Monetary Union. The empirical analysis is based on the theoretical equivalence relation that should hold between the CDS and bond spreads in a frictionless environment. We first test and find evidence in favour of the existence of persistent deviations between both spreads during the crisis but not before. Such deviations are found to be related to some market frictions, like counterparty risk, market illiquidity, and funding costs. We also find evidence suggesting that the price-discovery process is state-dependent. Specifically, the levels of counterparty and global risk, funding costs, market liquidity, volume of debt purchases by the European Central Bank in the secondary market, and the banks' willingness to accept losses on their holdings of Greek bonds are found to be significant factors in determining which market leads price discovery.

Keywords: sovereign credit default swaps, sovereign bonds, credit spreads, price discovery.

JEL Codes: G10, G14, G15.

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1 Introduction

In recent years many studies have analysed the relationship between credit default swaps (CDS) and bond spreads for corporate as well as for emerging sovereign reference entities.[1] However, the relation between sovereign CDS and bond markets in developed countries has not attracted much interest until very recently, mainly for two reasons. First, sovereign CDS and bond spreads in developed countries have been typically very low and stable given the perceived high credit quality of most issuers (see Table 1). Second, trading activity in this segment of the CDS market was typically low.

However, the global financial crisis that followed the collapse of Lehman Brothers in September 2008 triggered an unprecedented deterioration in public finances of the world's major advanced economies in a peacetime period. Since 2010, some countries in the euro area, including Greece, Ireland, and Portugal, and to a lesser extent Spain and Italy, have faced some episodes of heightened turbulence in their sovereign debt markets. Against this context, the levels of perceived credit risk and the volume of trading activity in the sovereign CDS markets in many advanced economies have increased.

The extant literature on credit risk has paid some attention to investigating the relationship between the corporate bond market and the corporate CDS market, but only a few papers have studied whether the empirical regularities identified in the corporate markets, including those related to price discovery, are also found in the case of sovereign reference entities. The aim of this paper is to shed light on these latter issues within the context of the recent episodes of sovereign-debt crises in several countries in the European Monetary Union (EMU).

Specifically, we analyse the theoretical equivalence relation between the sovereign bond yield spread (with respect to a risk-free benchmark) and the corresponding CDS spread.[2] Abstracting from market frictions and other contractual clauses, both spreads should reflect the same information on the credit risk of a given reference entity and therefore should be equal. In other words, the *basis*, defined as the difference between the CDS spread and the corresponding bond spread, should be zero. If the basis differs from zero, the differences should be purely random and unrelated to any systematic factor. Moreover, in such a frictionless scenario, both spreads (or credit-risk prices) should incorporate the credit-risk information in a similar way, i.e., both markets should be equally efficient in terms of the process of credit-risk price discovery. The current European sovereign debt crisis poses a particularly interesting scenario to test for the previous hypotheses. In particular, we analyse the bond-CDS equivalence relation from three different perspectives.

1 We discuss the related literature in Section 2.

2 The results are obtained using the German bond as a proxy of the risk-free asset, as in, e.g., Geyer et al. (2004), Bernoth et al. (2006), Delis and Mylonidis (2010), Favero et al. (2010), Foley-Fisher (2010), and Palladini and Portes (2011), among others.

First, we test the "no-arbitrage" theoretical frictionless relation that equates the bond and the CDS spreads. We find that there are persistent deviations from that relation. Interestingly, noticeable deviations begin with the outset of the subprime crisis, although no evidence of such deviations is found before this event.

Second, motivated by the previous finding, we study the possible causes of the deviations between the bonds and the CDS spreads. We find that the counterparty risk indicator has a negative and significant effect on the basis, especially after September 2008, when some of the most active protection sellers began to face financial difficulties. Funding costs have a negative effect on the basis due to their stronger effect on the demand for bonds relative to the demand for CDS, as the latter require less funding to take on the same risk position. A higher degree of liquidity in the bonds market relative to the CDS market has a positive effect on the basis given that *ceteris paribus,* a more liquid bond implies a lower bond yield and spread. The volume of debt purchased by the European Central Bank (ECB) in the secondary market that has taken place since May 2010 increases the basis significantly. These purchases exert a negative effect on bond spreads. The fact that such an effect is not present (or is weaker) in the case of the CDS spreads may indicate that ECB interventions affect other components of bond prices other than default risk (e.g., through a fall in the bond's liquidity premium) or, simply, induce some overpricing effect in the bond market for a given level of default risk. Although the effect of global risk, proxied by the VIX Index, is not significant, the country-specific risk premium, measured through the stock market index, affects the basis positively and significantly. This suggests that while global volatility is priced similarly in both markets, the idiosyncratic volatility is not, with the CDS market reacting more to changes in the latter case. Finally, the effect of the lagged basis suggests a high degree of persistence and, hence, a relatively low speed of adjustment of the basis.

Third, we address the question of which market leads the credit-risk price-discovery process. To this aim, we follow a dynamic price-discovery approach based on Gonzalo and Granger (1995). Our analysis reveals that the price-discovery process is state-dependent. Specifically, the levels of counterparty and global risk and the successive agreements of private banks to accept losses on their holdings of Greek bonds, impair the ability of the CDS market to lead the price-discovery process. The effect of counterparty risk is due to the perception of a lower quality of protection sold in the CDS market when this risk is high. The effect of global risk could be due to the fact that the information contained in bond spreads is more reliable during periods of high global risk. The agreements of private banks to accept losses on their holdings of Greek bonds could have caused a lack of confidence among investors in the CDS market after such agreements. On the other hand, the level of funding costs and the volume of sovereign debt purchased by the ECB worsens the efficiency of the bond market in the price-discovery process. Funding costs affect bond buyers more than they do CDS buyers, as the CDS market allows for more leveraged positions. The operations of the ECB seem to impair the informational content of bond prices as they relate to the actual credit risk of these assets.

The remainder of the paper is organised as follows: Section 2 discusses the related literature. Section 3 describes the data. Section 4 presents the methodology and the results based on the analysis of persistent deviations between CDS and bond spreads. Section 5 analyses the determinants of the basis. Section 6 presents the results of the dynamic price-discovery test. Section 7 contains some final remarks.

2 Related literature

In this section we focus on the branch of literature on CDS and bond spreads that is related to the three questions approached in this paper: persistent deviations between bond and CDS spreads, determinants of such deviations, and the price-discovery process in bonds and CDS markets.

We investigate the existence and persistency of deviations between CDS and bond spreads based on the notion of statistical arbitrage introduced by Hogan et al. (2004). As far as we know, this approach has only been applied to credit markets in the case of corporate CDS and bonds in Mayordomo et al. (2011a). In particular, they analyse the existence of persistent deviations between CDS and asset swap spreads of European corporations using the pre-crisis period (before 2008) and after the crisis period. Their results show that there are persistent deviations both in the pre-crisis and the crisis periods.

There is extensive literature addressing the determinants of corporate bond and CDS spreads.[3] Although this type of analysis is less frequent in the case of sovereign references, this topic is attracting increasing attention since the inception of the EMU.[4] Our aim, however, is not to study the determinants of the CDS or the bond spread, but, rather, the determinants of the basis to test whether both markets reflect different information. Although the analysis of the determinants of the basis is less frequent than the analysis of the individual credit spreads, there are some earlier contributions in the literature on sovereign credit markets. For instance, Fontana and Scheicher (2010) employ weekly data from 2006 to 2010 to analyse the determinants of the basis to find that the sovereign bases are significantly linked to the cost of short-selling bonds and to country-specific and global risk factors. In his analysis of CDS-bond parity, Levy (2009) finds that the frictionless parity relation does not hold for emerging markets' sovereign debt, but he argues that an important part of the deviations can be attributed to liquidity effects. Küçük (2010) relates the CDS-bond basis for 21 emerging market countries between 2004 and 2008 to factors capturing bond liquidity, speculation in CDS market, liquidity, equity market performance, and global macroeconomic variables. Foley-Fisher (2010) studies the relation between bond and CDS spreads for ten EMU countries on the basis of a theoretical model of heterogeneous investors' expectations. He shows that a non-zero basis is consistent with a relatively small

3 See, for instance, Elton et al. (2001), Collin-Dufresne et al. (2001), Chen et al. (2007), among others studying the determinants of the corporate bond spread. The studies analysing the determinants of the corporate CDS spreads include Longstaff et al. (2005), and Ericsson et al. (2009).

4 See, e.g., Codogno et al. (2003), Geyer et al. (2004), Bernoth et al. (2006), Favero et al. (2010), Beber et al. (2009), and Mayordomo et al. (2012).

dispersion in the beliefs of investors on the probability that certain European countries would default.[5]

We share some of the objectives pursued by these previous papers. However, to our knowledge, this work constitutes the first empirical analysis of the existence of persistent deviations in sovereign credit markets. Also, in contrast with previous analyses, our study of the determinants of the sovereign basis is carried out using daily data that includes the ongoing European Monetary Union sovereign debt crisis (May 2010-October 2011). The last scenario enables us to evaluate, among other factors, the effect of the purchases of sovereign debt by the ECB and the potential haircut on the banks' holdings of Greek bonds.

Finally, the most frequent analysis of the CDS-bond relation in corporate and sovereign credit markets is based on the concept of price discovery. Most recent papers study price discovery based on either Hasbrouck's (1995) or Gonzalo and Granger's (1995) methodologies. Both approaches build upon a test based on a Vector Auto Regression (VAR) with an Error Correction Term (ECT). For the period before the subprime crisis a recurrent empirical finding is that the CDS market reflects the information more accurately and quickly than the bond market in the corporate sector (see Blanco et al., 2005; or Zhu, 2006, among others). Most of the analyses of price discovery in sovereign markets have been applied to emerging markets. For instance, Ammer and Cai (2007) find that bond spreads lead CDS premia more often than what had been found for investment-grade corporate credits. Using data from eight emerging market countries for the period 2003-2006, Bowe et al. (2009) find that the CDS market does not, in general, lead price discovery, which appears to be country-dependent.

The recent crisis has renewed interest in this question in the context of the European sovereign debt markets. For instance, Fontana and Scheicher (2010) find that since the outset of the crisis, the bond market has had a predominant role in price discovery in Germany, France, the Netherlands, Austria, and Belgium, while the CDS market is playing a major role in Italy, Ireland, Spain, Greece, and Portugal. Palladini and Portes (2011) use data on six euro-area countries (Austria, Belgium, Greece, Ireland, Italy, and Portugal) over the period 2004-2011. They find that the CDS market moves ahead of the bond market in terms of price discovery for all the countries in the sample except for Greece. Delatte et al. (2010) find that the bond market leads the price-discovery process in the core European countries in periods of relative calm, while in periods of turbulence the CDS market leads the price-formation process. In the high-yield European countries, the CDS spreads reflect credit risk more adequately than the bond spreads in periods of both calm and tension, but the leadership of the CDS spread is exacerbated by financial turmoil. All these analyses have been carried out based on static measures of price discovery such that a single measure is obtained for the entire period analysed. However, as argued by Longstaff (2010), the price-discovery process in financial markets can be state-dependent. Thus, Delis and Mylonidis (2010) study the dynamic interrelation between bond and CDS spreads of several peripheral countries (Greece, Italy, Portugal, and

5 Analyses of the basis in the corporate credit market include Trapp (2009), Nashikkar et al. (2008), and Bai and Collin-Dufresne (2009), among others.

Spain) during the period July 2004 to May 2010 on the basis of a Granger causality test. They find bidirectional causality during periods of financial distress.

In the spirit of Longstaff's (2010) conjecture, we perform a state-dependent price-discovery analysis. Our paper estimates for the first time Gonzalo and Granger's (1995) price-discovery metrics over time. The use of this test allows us to overcome the bidirectional causality issue, which is commonly found by using the Granger causality test (see Delis and Mylonidis, 2010). We find methodological questions of the utmost importance given that determining which market leads at every period is essential to shed light on the factors that may influence the quality of a given market in terms of its power to contribute to the price-formation process. This paper also contributes to the previous literature by analysing a set of such factors in the context of the recent European sovereign debt crisis.

3 Data

The data consists of daily 5-year sovereign bond yields and CDS spreads for eleven EMU countries (Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, The Netherlands, Portugal, and Spain) from January 2004 to October 2011. Bond yields are obtained from Reuters, and CDS spreads from Credit Market Analysis (CMA), which reports data (bid, ask, and mid) sourced from 30 buy-side firms, including major global investment banks, hedge funds, and asset managers.

Table 1 reports the main properties of the data. As evident from this table, average CDS rates vary substantially across countries and periods. For the period 2004-2008, the lowest average CDS spread was 5 basis points (bp) for Germany and the highest one was 23 bp points for Greece. For the same period, the lowest average bond spread was 4 bp for both France and The Netherlands, and the highest average was 25 bp for Greece. For the period 2009-2011, the lowest annual average CDS spread was 31 bp for Finland in 2010 and the highest annual average was 2,075 bp for Greece in 2011 (being the maximum daily CDS spread at 6,752 bp on September 26th, 2011). The lowest annual average bond spread was -6 bp for Finland in 2010 and the highest was 1,644 bp for Greece in 2011.[6] We note that CDS spreads are on average higher than bond spreads in most of the countries, that is, the basis is positive (some of the most significant exceptions are Ireland and Portugal in 2011 and Greece in 2009 and 2010). Also, we observe an increase in both the average and the volatility of CDS and bond spreads over the subsequent years (from 2009 on) in most of the countries and especially in the peripheral ones (Greece, Ireland, Portugal, Spain, and Italy).

As for the rest of the data used in the subsequent estimations, the country-stock and global-risk indexes, which are proxied by means of the implied stock market volatility (we use the VIX for the global indicator), are obtained from Reuters. To capture funding costs we use the difference between the 90-day U.S. AA-rated commercial paper interest rates for financial companies and the 90-day U.S. T-bill, both from Datastream. We employ liquidity measures for the sovereign CDS and bonds, which are obtained from the bond and CDS bid-ask spreads. Bond bid-ask prices are obtained from Reuters, while CDS bid-ask spreads come from CMA. To proxy for the counterparty risk on the side of CDS dealers, we employ the CDS spreads of the 14 banks most active as dealers in the CDS market. These CDS spreads are obtained from CMA. The information regarding the European Central Bank (ECB) bond purchases, which took place after May 2010, was obtained from the ECB webpage.

6 The negative sign for the bond spread in Finland in 2010 is due to the fact that the average yield of the Finnish bond was lower than for the German bond.

CDS and bond spreads descriptive statistics

TABLE 1

			Bond	CDS
Austria	2004-2008	Mean	7	8
		Std. Dev.	12	21
	2009	Mean	52	104
		Std. Dev.	25	49
	2010	Mean	42	79
		Std. Dev.	12	13
	2011	Mean	59	90
		Std. Dev.	14	33
Belgium	2004-2008	Mean	9	8
		Std. Dev.	17	14
	2009	Mean	49	63
		Std. Dev.	29	33
	2010	Mean	57	110
		Std. Dev.	27	43
	2011	Mean	136	187
		Std. Dev.	47	51
Finland	2004-2008	Mean	5	6
		Std. Dev.	12	9
	2009	Mean	34	37
		Std. Dev.	21	19
	2010	Mean	-6	31
		Std. Dev.	13	3
	2011	Mean	18	44
		Std. Dev.	26	18
France	2004-2008	Mean	4	6
		Std. Dev.	9	9
	2009	Mean	23	40
		Std. Dev.	12	20
	2010	Mean	24	70
		Std. Dev.	8	18
	2011	Mean	41	107
		Std. Dev.	18	39
Germany	2004-2008	Mean		5
		Std. Dev.		7
	2009	Mean		36
		Std. Dev.		18
	2010	Mean		40
		Std. Dev.		8
	2011	Mean		58
		Std. Dev.		20
Greece	2004-2008	Mean	25	23
		Std. Dev.	36	36
	2009	Mean	166	165
		Std. Dev.	78	54
	2010	Mean	779	682
		Std. Dev.	290	242
	2011	Mean	1,644	2,075
		Std. Dev.	506	1,452

			Bond	CDS
The Netherlands	2004-2008	Mean	4	6
		Std. Dev.	8	12
	2009	Mean	30	53
		Std. Dev.	18	30
	2010	Mean	19	45
		Std. Dev.	7	8
	2011	Mean	24	55
		Std. Dev.	16	22
Ireland	2004-2008	Mean	9	13
		Std. Dev.	20	31
	2009	Mean	151	190
		Std. Dev.	56	62
	2010	Mean	262	302
		Std. Dev.	164	154
	2011	Mean	821	730
		Std. Dev.	235	145
Italy	2004-2008	Mean	19	19
		Std. Dev.	25	27
	2009	Mean	74	103
		Std. Dev.	32	39
	2010	Mean	109	165
		Std. Dev.	40	42
	2011	Mean	218	246
		Std. Dev.	101	115
Portugal	2004-2008	Mean	13	14
		Std. Dev.	20	19
	2009	Mean	71	76
		Std. Dev.	37	27
	2010	Mean	251	293
		Std. Dev.	111	116
	2011	Mean	918	767
		Std. Dev.	366	273
Spain	2004-2008	Mean	8	12
		Std. Dev.	15	20
	2009	Mean	54	89
		Std. Dev.	30	26
	2010	Mean	152	205
		Std. Dev.	74	67
	2011	Mean	257	295
		Std. Dev.	66	64

Table 1 reports the CDS and bond spreads main descriptive statistics (mean and standard deviation) for different time periods (2004-2008, 2009, 2010, and 2011). The bond spreads are obtained as the difference between country A's yield and the German yield.

4 Are there persistent deviations between CDS and bond spreads?

Suppose that an investor buys a bond at its par value with a maturity equal to T years and a yield-to-maturity equal to *ytm*. Also, assume that at the same time the investor buys protection on such reference entity for T years in the CDS market and the premium of such contract is s. The investor has eliminated the default risk associated with the underlying bond and the investor's net annual return is equal to $ytm - s$. Absent any friction, arbitrage forces would imply that the net return should be equal to the T-year risk-free rate, which we denote by r. Alternatively, if $ytm - s < r$, then by means of a short position in the bond, writing protection in the CDS market, and buying the risk-free bond the investor could have obtained a positive profit without any risk. If, on the contrary, $ytm - s > r$, the investor could obtain a certain profit by buying the risky bond, buying protection in the CDS market, and taking a short position in the risk-free bond. Hence, in equilibrium, $ytm - r = s$.

In order to investigate the existence and persistency of deviations between CDS and bond spreads that would violate the previous equilibrium relation, we apply the statistical arbitrage test employed by Mayordomo et al. (2011a). This test is based on the notion of arbitrage introduced by Hogan et al. (2004), according to which, absent market frictions, an arbitrage opportunity (in a statistical sense) represents a zero-cost, self-financing trading opportunity that has positive expected cumulative trading profits with a declining time-averaged variance and a probability of loss that converges to zero as time passes. Bearing in mind that within the logic of this methodology the existence of arbitrage opportunities is conditioned to the absence of market frictions, in our application of this test we interpret the results in a rather agnostic way. In particular, we do not identify persistent deviations between both spreads with unexploited arbitrage opportunities. Indeed, when such deviations are found we relate them, in a statistical sense, to several potential market frictions (see Section 5).

To test for the existence of persistent deviations from the zero-basis benchmark, we first compute the increase in the discounted trading profits that an investor would obtain under the assumption of no trading and funding costs. Specifically, the profits from a given investment strategy, in the sense just stated, are defined as the basis times the contract notional value. We compute such profits quarterly, and the payment on a given date t is added to the trading profits accumulated from the first investing date to the last date, t-1. The accumulated profits constructed in this way are assumed to have been invested or borrowed at the risk-free rate in the interim, from t-1 to t. The cumulative trading profits are then discounted up to the initial date. The increase in the discounted cumulative trading profits at a given date t is denoted by ffv$_t$ and is assumed to evolve according to the following process:

$$\Delta v_t = \mu t^{\theta} + \sigma t^{\lambda} z_t \tag{1}$$

for $t = 0, 1, 2, ..., n$, with n denoting the last investment date and z_t innovations. We assume the following initial conditions: $z_0 = 0$ and $v_0 = 0$ (i.e., the strategy is self-financed). Parameters θ and λ determine whether the expected trading profits and the volatility, respectively, are decreasing or increasing over time. Specifically, a positive θ (λ) indicates a time-increasing average (volatility) of the process; the higher this parameter, the stronger the speed of growth of the average (volatility) parameter. Under the assumption that z_t is an *i.i.d. N(0,1)* variable, the expectation and variance of the discounted incremental trading profits in equation (1) are $E\left[\Delta v_t\right] = \mu t^{\theta}$ *and* $Var\left[\Delta v_t\right] = \sigma^2 t^{2\lambda}$, respectively. Then, the discounted cumulative trading profits generated by a given strategy satisfy:

$$v_n = \sum_{t=0}^{n} \Delta v_t \sim N\left(\mu \sum_{t=0}^{n} t^{\theta}, \sigma^2 \sum_{t=0}^{n} t^{2\lambda}\right) . \tag{2}$$

We then define the log-likelihood function for the increments in equation (2) and estimate the parameters of interest $(\mu, \theta, \sigma, \lambda)$ by maximising that function using a non-linear optimisation method based on a Quasi-Newton-type algorithm. Then we formally implement the notion of the statistical arbitrage test outlined before through the specification and testing of the following three simultaneous hypotheses:

$$\begin{aligned}
&H1: \lim_{t\to\infty} E^P[v(t)] > 0 \quad \Rightarrow && \mu > 0, \quad and \\
&H2: \lim_{t\to\infty} P(v(t) < 0) = 0 \quad \Rightarrow && \lambda < 0 \quad or \quad \theta > \lambda, \quad and \\
&H3: \lim_{t\to\infty} Var\left[\Delta v(t) \mid \Delta v(t) < 0\right] = 0 \quad && \Rightarrow \theta > \max\left\{\lambda - \frac{1}{2}, -1\right\}.
\end{aligned}$$

Statistical arbitrage requires that the expected cumulative discounted profits*, v(t),* are positive (*H1*), the probability of loss converges to zero (*H2*), and the variance of the incremental trading profits *v(t)* also converges to zero (*H3*).[7]

Hence, these three conditions must be simultaneously satisfied to support the existence of persistent non-zero basis. In practice, this implies an intersection of several sub-hypotheses. To maximise the power of the test, instead of testing whether the previous hypotheses are simultaneously satisfied, we redefine the null hypothesis as the absence of persistent non-zero basis, and so our test is based on the following union of sub-hypotheses, which are given by the complementarity of the previous hypotheses (see Jarrow et al., 2011):

$$\begin{aligned}
&H1^C : \mu \le 0, \quad or \\
&H2^C : \lambda \ge 0 \quad and \quad \theta - \lambda \le 0, \quad or \\
&H3_1^C : \theta - \lambda + \frac{1}{2} \le 0, \quad or \\
&H3_2^C : \theta + 1 \le 0,
\end{aligned}$$

7 Implicit in hypothesis *H3* is the idea that investors are only concerned about the variance of a potential decrease in wealth. Whenever the incremental trading profits are non-negative, their variability is not penalised.

where $H1^C$ and $H2^C$ are the complementaries of hypotheses *H1* and *H2*, respectively, while $H3_1^C$ and $H3_2^C$ come from the complementary of hypothesis *H3*. If at least one of the last four hypotheses above is satisfied, we conclude that no persistent deviations exist.

To test these hypotheses we need to estimate the *p-values* for the previous restrictions. To this aim, we follow the methodology developed by Politis et al. (1997, 1999). This technique provides an asymptotically valid test under weak assumptions. Specifically, our analysis leads to two one-tail tests:

a) H_o: no persistent deviations and H_A: negative deviations (the bond spread is significantly higher than the CDS spread);

b) H_o: no persistent deviations and H_A: positive deviations (the CDS spread is significantly higher than the bond spread).

The results of these tests are summarised in Table 2. Panels A and B report the results for the period ranging from January 2004 to September 2008 for negative and positive bases, respectively. Panels C and D report the corresponding results for the period ranging from September 2008 to October 2011. As shown in Panels A and B, we cannot reject the null hypothesis (no persistent deviations) at any standard significance level. This result holds irrespective of whether we consider either positive or negative bases. However, after September 2008 the CDS spread is persistently higher than the bond spread in six cases (see Panel D), while none of the countries analysed presents a persistent negative basis, shown in Panel C.

As a conclusion, the above results reveal that the zero-basis hypothesis cannot be rejected when we consider the pre-crisis period, although temporary non-zero bases are not rare during the crisis. This last result must be interpreted with caution since, as argued before, a non-zero basis cannot be understood mechanically as an opportunity for arbitrage. For instance, Schonbucher (2003) and Mengle (2007) emphasise that shorting a bond with a required maturity is not always feasible. Moreover, the fact that recurrent non-zero bases seem to be common during the crisis period may be symptomatic of the presence of other restrictions and frictions that prevent a perfect timeless alignment between the CDS and the bond spreads and whose relevance may have been exacerbated by the crisis itself. This could be the case, for instance, of funding costs, differences in liquidity across markets, and counterparty risk in the CDS market. In the following section we test for the significance of these (and other) factors as potential explanatory variables for the cases of non-zero basis detected during the crisis.

Statistical arbitrage test for the existence of persistent mispricings TABLE 2

	P-value	Persistent mispricing
Panel A: Persistent negative basis before Lehman Brothers collapse		
Austria	1.000	No
Belgium	0.961	No
Finland	1.000	No
France	1.000	No
Greece	0.999	No
The Netherlands	0.988	No
Ireland	1.000	No
Italy	0.678	No
Portugal	1.000	No
Spain	0.988	No
Panel B: Persistent positive basis before Lehman Brothers collapse		
Austria	1.000	No
Belgium	0.957	No
Finland	1.000	No
France	1.000	No
Greece	1.000	No
The Netherlands	0.987	No
Ireland	0.706	No
Italy	0.378	No
Portugal	1.000	No
Spain	0.988	No
Panel C: Persistent negative basis after Lehman Brothers collapse		
Austria	1.000	No
Belgium	1.000	No
Finland	1.000	No
France	1.000	No
Greece	0.641	No
The Netherlands	1.000	No
Ireland	1.000	No
Italy	0.650	No
Portugal	0.734	No
Spain	0.753	No
Panel D: Persistent positive basis after Lehman Brothers collapse		
Austria	0.016	Yes**
Belgium	1.000	No
Finland	0.963	No
France	0.043	Yes**
Greece	0.968	No
The Netherlands	0.003	Yes***
Ireland	0.047	Yes**
Italy	0.034	Yes**
Portugal	1.000	No
Spain	0.041	Yes**

This table reports the p-value obtained from the statistical arbitrage methodology of Mayordomo et al. (2011a). A p-value lower than 0.05 indicates that at a significance level of 5% there are persistent mispricings between the 5-year CDS and bond spreads. The bond spread is obtained as the difference between country A's bond yield and the risk-free rate, which is equal to the German bond yield. Panels A and B report the results for the period ranging from January 2004 to September 2008 for CDS-bond negative and positive bases, respectively. Panels C and D report the results for the period that spans from the collapse of Lehman Brothers (September 2008) to October 2011 for CDS-bond negative and positive bases, respectively. *** (** and *) indicates the existence of persistent mispricings at a significance level of 1% (5% and 10%, respectively).

5 The determinants of the basis

In this section we test whether the differences between the CDS and bond spreads are purely random or, alternatively, whether they are related to any market-specific or global factors. Due to observations corresponding to the last months of 2010 and the months of 2011 included in the sample, the basis behaves like a non-stationary variable. For this reason, instead of analysing the determinants of the basis we study the determinants of the relative basis, defined as the difference between the CDS and bond spreads relative to the average credit spread, which is obtained as the simple mean of the CDS and the bond spreads. We consider the following potential explanatory factors:

a. **Counterparty Risk**. In principle, the higher the counterparty risk of the seller of a CDS, the lower the CDS price should be as a result of the lower quality of protection. We test for this effect by using the first principal component obtained from the CDS spreads of the main 14 banks that act as dealers in that market.[8] The first principal component series should reflect the common default probability and, hence, it is akin to an aggregate measure of counterparty risk.[9] Actually, the first principal component for the series of CDS spreads of this set of dealers explains 87.5% of the total variance of the observed variables. We use the counterparty risk variable lagged one period to avoid problems of endogeneity derived from the potential contemporaneous effects of the banks' activity on sovereign credit spreads.

b. **Liquidity.** In theory, one would expect that higher liquidity in the bond market relative to the CDS market would go hand in hand with a higher basis, since a more liquid bond implies a lower spread in that market. To test for relative liquidity effects, we construct a ratio of relative liquidity between the CDS and the bond. Specifically, the degree of liquidity in the CDS market is proxied by the bid-ask spread of the CDS premium. The higher this spread is, the lower the degree of liquidity in the CDS market. A similar measure of liquidity is computed for the bond market. The ratio between both measures is taken as indicative of the relative liquidity in the CDS market vis-à-vis the bond market. As this ratio rises, liquidity in the CDS market relative to the bond market falls. Therefore the basis would, in principle, increase.

8 The 14 main dealers are: Bank of America, Barclays, BNP Paribas, Citigroup, Credit Suisse, Deutsche Bank, Goldman Sachs, HSBC, JP Morgan, Morgan Stanley, Royal Bank of Scotland, Societé Generale, UBS, and Wachovia/Wells Fargo. These dealers are the most active global derivatives dealers and are known as the G14 (see, for instance, ISDA Research Notes, 2010, on the Concentration of OTC Derivatives among Major Dealers).

9 The use of the dealers' CDS spreads as a proxy of counterparty risk is based on the Arora et al. (2009) study, which analyses the existence of counterparty risk in the corporate CDS market.

c. **Financing Costs**. One would expect that higher financing costs would lower the demand for bonds and could lead to a decrease in prices, and therefore to higher bond spreads. The effect of funding costs on CDS spreads should be weaker given that in this case the required amount of funding to get the same (gross) risk position is lower (i.e., risk-leverage is higher in the case of the CDS investment). For this reason, in principle, an increase in financing costs would have a negative effect on the basis. Due to the difficulty in obtaining data on institution-level funding constraints, we use the spread between financial commercial paper and T-bill rates as a common proxy for the funding constraints faced by financial intermediaries, as in Acharya et al. (2007). Specifically, we use the spread between the 90-day U.S. AA-rated commercial paper interest rates for the financial companies and the 90-day U.S. T-bill.

d. **Domestic and Global Risk Premiums.** As additional potential explanatory variables for the basis, we consider a measure of the country and global risk premium. If both the CDS and bond spreads are prices for the same credit risk, the effect of the country-specific and global-risk premia on the basis should not significantly differ from zero. However, a significant non-zero effect would suggest that one of the markets reacts more to changes in the risk premiums than the other. To control for the fact that this idiosyncratic and global volatility could be priced differently in the two markets, we use the previous risk factors as additional explanatory variables. The country-specific risk premium is proxied by the stock market volatility, defined as the absolute value of the stock index returns. The global risk premium or global risk is proxied by the VIX Index. The correlation between the VIX Index and the counterparty risk variable is around 0.8. Thus, in order to avoid any multicollinearity problem, we modify the VIX Index variable and define it as the residual of the regression of this variable onto the first principal component CDS spreads corresponding to the main CDS dealers such that counterparty risk and the VIX are now orthogonal variables.

e. **Volume of Debt Purchased by the European Central Bank.** We use as an additional explanatory variable the amount of sovereign debt purchased by the ECB in the secondary market from May 2010 onwards. This information is available only on a weekly basis, and so the effect of the ECB debt purchases is proxied by the amount of debt that was purchased the week before the current date. These purchases are supposed to decrease both CDS and bond spreads, but in principle it is natural to think that they would have a more direct effect on bond prices. Thus, following these interventions, the bond spread would decrease more than the decrease in the CDS spread, thus affecting the basis positively.

f. **Dummy for the Private Banks' Initial Pre-Agreement to Accept Losses on Their Holdings of Greek Bonds.** The proposal launched in July 2011 for an agreement among banks to accept a voluntary haircut on their holdings of Greek bonds, which could come without the activation of the CDS contracts, could have generated a lack of confidence among institutional investors in the CDS market. This last effect may lead to a lower quality of the protection sold, which should imply lower CDS spreads and a lower basis. For this reason, we use as an additional regressor a dummy variable that is equal to 1 after the banks agreed on the 21st of July 2011 to accept 21% losses on their holdings of Greek bonds and zero before that date.

g. **Lagged Basis.** The lag of the basis should absorb any past information transmitted into the current observation. Due to the existence of persistent deviations between the CDS and bond spreads documented in Section 4, we expect a positive sign.

We estimate the coefficients for the above factors for the period spanning January 2004 to October 2011 using a fixed-effects estimation procedure that is robust to heteroskedasticity. We use the bootstrap methodology to correct for any potential bias in the standard errors due to the use of generated regressors. The results are reported in Table 3. Column 2 of that table reports the standardised coefficients (i.e., the regression coefficient as in Column 1 multiplied by one standard deviation of the corresponding explanatory variable). All the variables, including the dependent variable, are expressed on a per unit basis, with the exception of the volume of bonds purchased by the ECB in a given week, which is defined in billions of euros.

The counterparty risk proxy has a negative and significant effect, as expected, which confirms its relevance on the levels of the CDS spreads. Funding costs have a negative effect due to their stronger effect on the demand for bonds relative to the demand for CDS that requires a lower amount of funding. A low degree of liquidity in the CDS market relative to the bond market has a positive effect given that a more liquid bond implies a lower bond spread. The global-risk variable is not significant, which may suggest that both markets reflect global risk to a similar extent. Nevertheless, the country-specific risk measure seems to be priced differently in both markets. The positive and significant effect could be explained by the fact that the CDS spread reacts more to the variation in aggregate, macro factors, such as country-specific risk premium, than the bond spreads, a feature that has been documented by Alexopoulou et al. (2009) and Mayordomo and Peña (2011). The dummy for the banks' willingness to accept losses on their holdings of Greek bonds has a negative sign, as expected, but it is not significant at any standard significance level. In line with the results of the previous section, we find a high level of persistency in the relative basis. That is, there is a relatively low speed of adjustment towards the long-run bond-CDS equivalence relation. Finally, the constant term reflects whether the relative basis differs, on average, from zero and the magnitude of such deviation. We observe that the relative basis is on average significantly positive, suggesting that the bond-CDS equivalence relation does not hold even when we take into account the market frictions described above and the costs that are needed to trade the basis. Nevertheless, its magnitude is low regarding the average relative basis during the period 2004-2011 (2% relative to 32%). Thus, when we take into account the determinants of the basis, the magnitude of the average relative deviation is close to zero and is reduced by 93.75%. This result confirms the strong influence on these determinants to guarantee the bond-CDS equivalence relation.

The relatively high R-square of this regression is mainly due to the effect of the lagged basis. However, it should be noted that the explanatory variables retain a relatively high explanatory power even when we ignore the lagged basis, in which case the R-square is 0.29. Actually, this is of a similar magnitude to the one reported by Trapp (2009) on a daily basis for corporates using firm fixed effects but ignoring the effect of the lagged basis.

In Column 2 we report the standardised coefficients. As expected, the strongest effect is the one corresponding to the lagged relative basis. Among the rest of the variables, the strongest effects are the ones coming from the counterparty risk and the financing costs. An increase equal to one standard deviation in any of the previous variables would diminish the relative basis in –0.7%.

Determinants of the basis TABLE 3

	(1)	(2)
Counterparty risk	-0.163***	-0.005
	(0.06)	
Ratio CDS/bond liquidity	0.004***	0.003
	(0.00)	
Financing costs	-1.596***	-0.008
	(0.31)	
Global risk (VIX) net of counterparty risk	-0.031	-0.002
	(0.02)	
Volatility of country stock index returns	0.398***	0.004
	(0.12)	
Bonds purchased by ECB	0.001***	0.002
	(0.00)	
Dummy haircut agreement by banks (July 11)	-0.002	0.000
	(0.00)	
Lagged relative basis	0.933***	0.566
	(0.01)	
Constant	0.022***	
	(0.00)	
Number of observations	16,127	
Wald Chi2 statistic	46,367	
Prob > Wald Chi2	0	
Adjusted R-squared	0.905	

This table reports the effect of the potential determinants of the basis based on a fixed-effects regression robust to heteroskedasticity. Column (1) reports the effect of such determinants for the period spanning January 2004-October 2011. This column contains the explanatory variables' coefficients and the standard errors between brackets. *** (** and *) indicates whether the coefficients are significant at a significance level of 1% (5% and 10%). The bootstrap methodology is employed to correct any potential bias in the standard errors due to the use of generated regressors. Column 2 reports the standardised coefficient (i.e., the regression coefficient as in Column 1 multiplied by the standard deviation of the corresponding explanatory variable).

6 Price-discovery analysis

An efficient price-discovery process is characterised by a quick adjustment of market prices from the old to the new equilibrium as new information arrives (see, e.g., Yan and Zivot, 2007). In the analysis that follows, we show that the price-discovery process in the markets for sovereign credit risk in the euro area shows a time-variant pattern (Section 6.1). We then try to identify the effect of several potential explanatory variables of the dynamic pattern followed by the price-discovery metrics obtained in the previous step (Section 6.2).

6.1 A dynamic price-discovery metric

To estimate a time-variant price-discovery metric we extend Gonzalo and Granger's (1995) price-discovery analysis using rolling windows. Gonzalo and Granger's model of price discovery is based on the following vector error correction model (VECM) specification:

$$\begin{pmatrix} \Delta BSpr_t \\ \Delta CDSSpr_t \end{pmatrix} = \alpha(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1}) + \begin{pmatrix} \sum_{i=1}^{p} \lambda_{1,i} \Delta BSpr_{t-i} \\ \sum_{i=1}^{p} \lambda_{2,i} \Delta BSSpr_{t-i} \end{pmatrix} + \\ + \begin{pmatrix} \sum_{i=1}^{p} \delta_{1,i} \Delta CDSpr_{t-i} \\ \sum_{i=1}^{p} \delta_{2,i} \Delta CDSSpr_{t-i} \end{pmatrix} + \begin{pmatrix} u_{1,t} \\ u_{2,t} \end{pmatrix}. \tag{3}$$

The above empirical model is a vector autoregressive (VAR) system formed by two equations defined from the vector, which includes the bond and CDS spreads of the same underlying country, denoted by $BSpr_t$ and $CDSSpr_t$, respectively, and an error correction term (ECT) defined by the expression $\alpha(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1})$, where β_2 and β_3 are estimated in an auxiliary cointegration regression and the parameter vector $\alpha' = (\alpha_1, \alpha_2)$ contains the error-correction coefficients measuring each price's expected speed of adjustment. The estimation of the VECM specification is restricted to the existence of a cointegration relation between the bond and CDS spreads. This cointegration relation appears in the ECT as $\left(BSpr_{t-1} - \beta_2 - \beta_3 CDSSpr_{t-1}\right)$. The parameters $\lambda_{1,i}$, $\lambda_{2,i}$, $\delta_{1,i}$, and $\delta_{2,i}$ for $i = 1,\dots p$, with p indicating the total number of lags, contain the coefficients of the VAR system that measure the effect of the lagged first difference in the CDS and bond spreads on the first difference of such spreads at time t.[10] Finally, u_t denotes a white noise vector.

10 The optimal number of lags is determined by the Schwarz information criteria.

The price-discovery metric for the bond and CDS markets, denoted by GG_{bond} and GG_{CDS}, respectively, can then be constructed from the elements of the vector α', which contains the coefficients that determine each market's contribution to price discovery:

$$GG_{Bond} = \frac{\alpha_2}{-\alpha_1 + \alpha_2}; \qquad GG_{CDS} = \frac{-\alpha_1}{-\alpha_1 + \alpha_2} \quad . \tag{4}$$

Given that $GG_{Bond} + GG_{CDS} = 1$, we would conclude that the bond (CDS) market leads the price-discovery process whenever GG_{Bond} is higher (lower) than 0.5. The intuition for this is the faster the speed in eliminating the price difference from the long-term equilibrium attributable to a given market, the higher the corresponding α according to (3), and the higher the price-discovery metric.

In order to apply the methodology outline above to a dynamic metric of price-discovery leadership in the two markets at stake, we estimate the system in equation (3) using rolling windows with different lengths: 500, 750, and 1,000 days. To do so, we first need to check for the order of integration of the CDS and bond spreads and then for the existence of a cointegration relation. An interesting feature is that, on average, the estimated metrics do not seem to be very sensitive to the window length. Based on this, in the remainder of the paper we focus on 1,000-day windows.[11] Using rolling windows with a length of 1,000 observations, we find that both CDS and bond spreads are non-stationary in all the countries and dates.[12] We next apply the cointegration test to a total of 990 1,000-day windows for each of the ten countries to find cointegration between both spreads in 6,295 cases (64% of the total). The country with the lowest (highest) percentage of cointegration relations is Finland (Portugal) with cointegration in 41% (83%) of the windows. As we increase the window length, we find a higher number of cointegration relations and unit roots.

Figure 1 shows the estimated price-discovery metric for the 1,000-day windows for two groups of countries in the sample, peripheral and central.[13] In particular, we report a 30-day moving average of the mean price-discovery metrics, which is obtained as an equally weighted mean across the ten euro-area countries. An important message in Figure 1 is that the price-discovery metrics are not static but rather evolve over time, with the relative leadership of the CDS market in the process of price discovery being more pronounced around specific dates.

Specifically, before the Lehman Brothers collapse the CDS market led sovereign risk price discovery. This finding is consistent with the results reported by, for example,

11 On the one hand, the 500-day windows enable us to update the influence of the new observations quicker than the 1,000-day windows and to estimate the price-discovery metrics for the years 2006 and 2007. On the other hand, the 1,000-day windows enable us to consider a higher number of price-discovery metrics per day and country after 2007 due to the existence of a higher percentage of cointegration relations for these windows. Actually, the years 2006 and 2007 present the lowest number of days with cointegration relations per year.

12 The number of lags employed in the unit root test is chosen according to the Schwarz information criterion.

13 The peripheral group includes Ireland, Italy, Greece, Portugal, and Spain. The core group includes Austria, Belgium, Finland, France, and The Netherlands. The window length in all cases is 500 days.

Blanco et al. (2005) or Zhu (2006) in the context of the corporate debt markets. The first noticeable rise in the relative leadership of the bond market took place around February 2008, around the collapse of Bear Stearns. Afterwards, in September 2008, coinciding with the fall of Lehman Brothers and AIG, the bond price-discovery metric again jumped reaching its highest value at the end of 2008. This pattern suggests that during these two specific episodes, the collapse of Bear Stearns and Lehman-AIG, the bond spread led, by a small margin, the price-discovery process. The price-discovery metrics show the last significant rebound around the end of June 2011, some days before the proposal for an agreement among some private banks in July 2011 to accept a loss of 21% on their Greek bonds under the implicit presumption that such a voluntary deal would not trigger a credit event that would call for the activation of CDS protection. Following this, in August 2011 the bond market became the leader in terms of price discovery, confirming the role of this market during the most recent phase of the crisis as the fairest source of information on credit risk.[14]

From a different perspective, we observe a decoupling of the price-discovery measure for the core and peripheral countries from the end of 2008 until mid-2011. Starting at the end of 2008, it is worth noting that for most of the time the relative efficiency of the CDS market in the peripheral countries is significantly higher than for the core countries, where the bond market indeed led the price-discovery process for most of 2009. The difference between the estimated price-discovery metric for both groups of countries widened further during 2010. This pattern, which is mainly motivated by a sharp increase in the relative efficiency of the CDS market, may reflect the ECB policy of buying sovereign debt issued by peripheral countries.

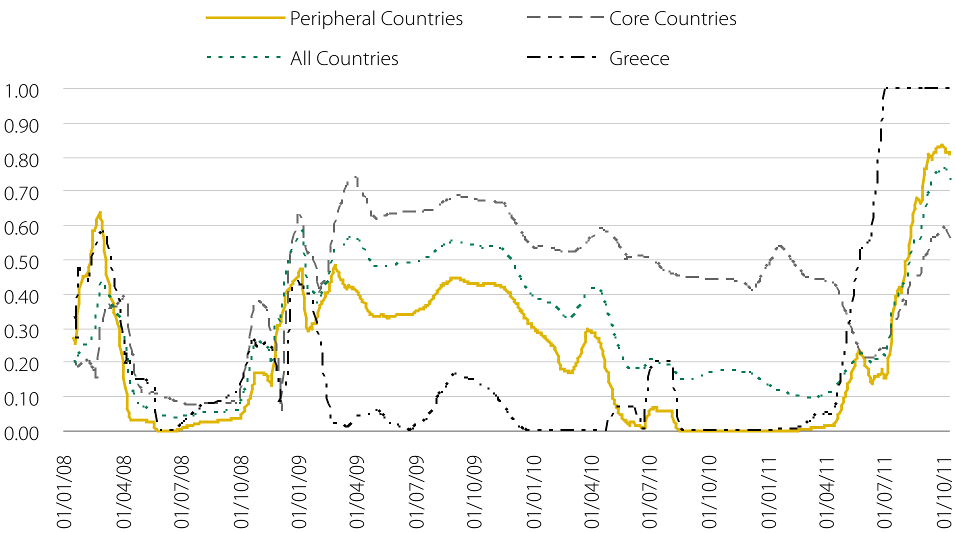
This core-peripheral countries decoupling persisted until May 2011 when we observed a similar trend in both groups of countries of improving the power of the bond market to lead price discovery. In particular, after July 2011 the bond market moved ahead of the CDS market. This result could be due to the increase of overall risk perception, but it also could be due to a lack of confidence of institutional investors in the CDS market after the aforementioned proposal for an agreement among the banks to accept a voluntary haircut on their Greek bonds without activating the CDS contracts. This explanation would fit well with the fact that the price-discovery metric for Greece jumps from a value near zero in April 2011 to 1 in July 2011.

Figure 2 shows the estimated price-discovery metrics for the 1,000-day windows for the same two groups of countries (peripheral and core), but it shows the number of countries for which we can implement the Gonzalo and Granger methodology (i.e., unit roots in both the CDS and bond spreads series are found and both spreads are cointegrated). The darker the line is, the higher the number of countries for which the analysis can be implemented. We observe that for most of the time, the analysis can be implemented in four or more countries. Actually, it can be implemented in eight or more countries for long windows of time from the beginning of 2008 to the end of 2009.

14 This result is in line with the role of the bond market as an appropriate credit-risk measure in crisis periods found by Mayordomo et al. (2011b) for the corporate case.

Price-discovery metrics for groups of EMU countries with 1,000-day rolling windows

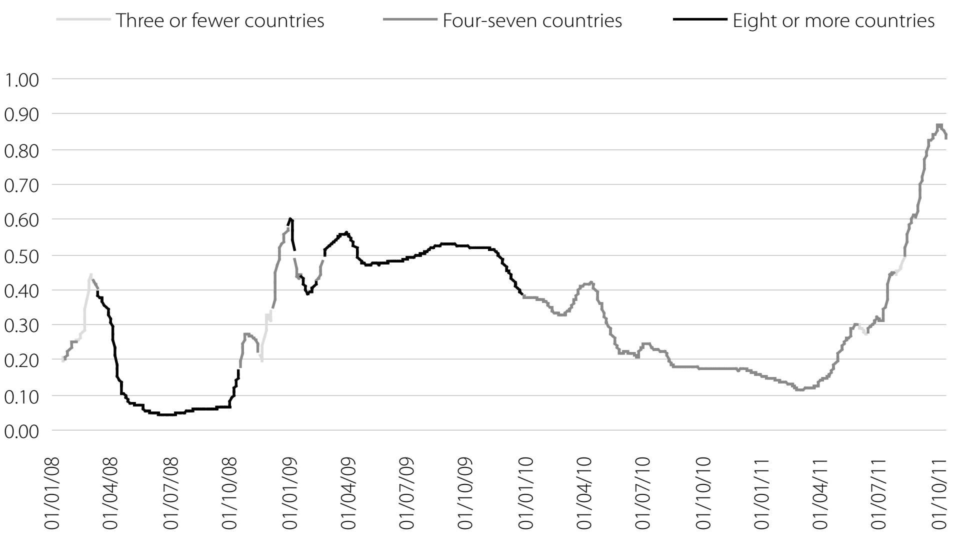
FIGURE 1



This figure shows the 30-day moving average of the price-discovery metrics for the peripheral, core, and all EMU countries. The price-discovery metrics are estimated using 1,000-day rolling windows. These metrics for the groups of countries are obtained as the equally weighted average of the country-specific price-discovery metrics. The price-discovery metrics for Greece correspond to the 30-day moving average of the metrics for Greece.

EMU price-discovery metrics and number of countries employed in their calculation

FIGURE 2



This figure shows the 30-day moving average of the EMU countries' price-discovery metrics that are obtained using 1,000-day rolling windows. These metrics are obtained as the equally weighted average of the country-specific price-discovery metrics. The line shows the number of countries employed to calculate the average metric such that the darker the line, the higher the number of countries employed in its calculation.

6.2 An analysis of the determinants of market leadership in price discovery

In this section we aim to shed some light on the dynamic pattern of the price-discovery metrics estimated before by regressing them on potential explanatory factors. Specifically, for each country we construct a dummy variable that takes a value of 1 when the bond market reflects information more efficiently than the CDS market and zero otherwise. This dummy is constructed on the basis of a rolling window estimation using 1,000 observations,[15] and then it is used as the dependent variable in a Logit regression that includes as regressors the same used in the regression contained in Table 2, with the exception of the lagged basis. We consider these regressors because they have been found to have a significant effect on the deviations from a zero basis. Hence, as this shows that the effect of such regressors is not reflected in the same way in the two markets, it seems natural to consider that one market captures better than the other the effect of each determinant of the basis.

The results are reported in Table 4. Columns 1 and 3 report the results obtained when we use daily and monthly price discovery metrics, respectively, for the period spanning December 2007-October 2011. Column 2 contains the marginal effects of the coefficients reported in Column 1. The marginal effects of the dummy variables are obtained as the discrete change (from 0 to 1) in the corresponding variable. All the explanatory variables, including the dependent variable, are expressed in percentages with the exception of the bonds purchased by the ECB in a given week, defined in billions of euros, and the dummy haircut agreement by banks, which is equal to 1 after the 21st of July 2011 and zero before then.

One might expect that the higher the counterparty risk, the lower the power of the CDS market to adequately reflect the credit risk due to the lower quality of the protection sold in this market and, perhaps, the uncertainty around such quality. The sign is the expected (positive) and significant. An increase equal to one percentage point in the counterparty risk measure would increase the contribution of the bond market to the leadership of the price discovery process by 3.5%.

As argued before, funding costs affect the bond market relative to the CDS more negatively, as the CDS market allows for higher-leveraged positions. This could explain why the funding costs negatively affect the ability of the bond market to anticipate the price of credit risk relative to the CDS market. Specifically, we find that an increase of 1% in financing cost would decrease the estimated price-discovery metric by 23.6%. Surprisingly, the degree of liquidity in the CDS market relative to the bond market does not significantly affect the price-discovery metric.

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15 When faced with missing values (due to a lack of cointegration relation between the CDS and the bond spreads), the value of the dummy is imputed according to the Granger-causality analysis performed earlier. For the remaining missing values, the value of the dummy is imputed whenever there is another observed value within the next ten days that coincides with the previous price discovery value observed and persists at least in the next ten observations. If after a given date there are not cointegration relations up to the end of the sample we do not impute any value. For the sake of the robustness of this procedure, we use the 1,000-observation estimation. Thus, we do not employ the price-discovery metric directly, which is a concrete value between 0 and 1, but instead assign a value of 1 or 0 to such metrics, thus softening such a strong assumption. We impute 16% of the total observations.

In line with the results obtained by Mayordomo et al. (2011b), the bond spreads tend to reflect credit risk more efficiently than CDS spreads during periods of high global risk (high values of the VIX Index). On the contrary, the country-specific risk premium proxied by the absolute value of the returns of the national stock index is not significant at any standard significance level.

If the ECB's demand for debt is relatively inelastic with respect to its price then the information embedded in the prices formed in that market could reveal less about the fundamental value of the corresponding bonds. This hypothesis is confirmed by the significant and negative sign of the variable representing the total amount of sovereign debt purchased by the ECB.[16] An increase of $1 bn in the total amount of debt purchased by the ECB leads to a decrease of -0.3% in the ability of the bond market in terms of price discovery. The total amount purchased by the ECB at the end of the sample period (October 2011) was $180.5 bn. Such an increase could have contributed to a 54% fall in the estimated price-discovery metric, thus leading to a large deterioration in the quality of price information in the bond market.

The proposal for an agreement among private banks in July 2011 to accept a voluntary loss on their holdings of Greek bonds seemed to favour the relative efficiency of the bond markets, according to the positive and significant effect of the corresponding dummy variable. Specifically, after the 21st of July 2011, we find a fall of 17.8% in the contribution of the CDS market to the price-discovery process.

Finally, we note that the use of daily price-discovery metrics obtained with 1,000-day rolling windows implies that the new information added in a given day is small relative to the information of the other 999 days, which persists from one estimation to the next. Faced with this, we check whether the results hold independently or whether we use daily or monthly price-discovery metrics. The corresponding results are reported in Column 3. The comparison of the results reported in Columns 1 and 3 confirms that the results are robust regarding daily or monthly metrics.

16 Given that we are using historical information on bond and CDS spreads to construct the price-discovery metric, we now use the stock of debt purchased by the ECB instead of the flow corresponding to the previous week.

Determinants of the price-discovery metrics

TABLE 4

	(1)	(2)	(3)
Counterparty risk	0.287***	0.035	0.263***
	(0.07)		(0.10)
Ratio CDS/bond liquidity	0.000	0.000	-0.003
	(0.00)		(0.00)
Financing costs	-1.932***	-0.236	-1.897***
	(0.53)		(0.48)
Global risk (VIX) net of counterparty risk	0.050**	0.006	0.092**
	(0.02)		(0.04)
Volatility of country stock index returns	-0.001	0.000	-0.251
	(0.05)		(0.44)
Bonds purchased by ECB	-0.024***	-0.003	-0.012**
	(0.01)		(0.01)
Dummy haircut agreement by banks (July 11)	1.453**	0.178	2.452***
	(0.73)		(0.84)
Constant	-1.987***		-1.223
	(0.35)		(0.82)
Number of observations	8,295		379
LR Chi2 statistic	3,750		101
Prob > Wald Chi2	0		0
Log pseudolikelihood	-2,888		-142
Pseudo R2	0.394		0.398

This table reports the effect of the potential determinants of the price-discovery metrics using a panel fixed--effects logistic regression robust to heteroskedasticity. The standard errors are clustered by country. The price-discovery metrics are obtained from Gonzalo and Granger's (1995) methodology using rolling windows of 1,000 observations. The dependent variable takes a value of 1 when the bond spread reflects the information more efficiently than the CDS spread, while a value equal to 0 indicates that the CDS spread leads the price discovery process. The bond spread is defined as the difference between country A's yield and the German yield. Column (1) reports the results obtained when we use daily price-discovery metrics for the period spanning December 2007-October 2011. Column (2) includes the marginal effects of the coefficients in Column (1). Column (3) reports the results obtained when we use monthly price-discovery metrics for the same period. The sample length is due to the use of the first 1,000 observations to estimate the price-discovery metric. The table contains the explanatory variables' coefficients and the standard errors between brackets. *** (** and *) indicates whether the coefficients are significant at a significance level of 1% (5% and 10%).

7 Conclusions

This paper analyses the extent to which the sovereign CDS and bond markets reflect the same information on their prices in the context of the European Monetary Union. The main results can be summarised in the following.

We first test the "no-arbitrage" theoretical relation that should exist between the bond and the CDS spreads in a frictionless environment since both spreads are supposed to be the prices for the same credit risk. Our results show that after the subprime crisis there were persistent deviations from the theoretical parity relation that were absent before. In particular, we find evidence in favour of a persistent positive basis for the crisis period in a number of countries.

Based on the previous finding, we analyse the role of some potential determinants of the basis, including several sources of risk (counterparty, country-idiosyncratic, and global) and market frictions. In particular, we find that the counterparty risk indicator has a negative and significant effect on the basis. Funding costs have a negative effect on the basis, while a high liquidity in the bond market relative to the CDS market has a positive effect. Although the effect of the global-risk variable is not significant, the country-specific risk premium measured by means of the stock market index affects positively to the basis. The European Central Bank's purchases of sovereign debt in the secondary market in May 2010 were found to positively and significantly affect the basis.

Finally, we conduct a dynamic analysis of market leadership in the price-discovery process. An important result here is that the price-discovery process is clearly state-dependent. Specifically, the levels of counterparty and global risk, funding costs, the volume of debt purchases by the European Central Bank following the rescue of the Greek economy, and the subsequent proposal for an agreement among private banks to voluntarily accept losses of their Greek bonds are all found to be significant factors in determining which market leads price discovery.

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